Assessing the Forecasting Performance of Structural Models for the Nominal Exchange Rate: The Colombian Case*.

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Abstract:

This paper analyses the out-of-sample forecasting performance of three models for the USD/COP nominal exchange rate during the period 1984:I – 2004:I. The sticky price monetary (Dornbusch (1976) – Frankel (1979)) and the Balassa–Samuelson (which gives a central role to the productivity differentials) approaches are used. Additionally, the Purchasing Power Parity condition (PPP) is analyzed. The forecasting ability of these models is compared using a random walk as a benchmark model. The metrics employed in evaluating the forecasting performance are RMS, MAE, MAPE and U-Theil. It is found that despite the great ability to predict, no model outperforms the random walk. This conclusion strengthens the previous results in the nominal exchange rate modeling literature.

Keywords: Nominal Exchange Rate, Econometric Models, Colombia, Forecasts
JEL Classification: F310, C220, C530

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I. INTRODUCTION

Movements in the USD/COP nominal exchange rate[1] during the last ten years and especially during the year 2004 have left analysts, policymakers and exchange market agents disturbed. The low ability to predict the exchange rate movements appears to reveal a huge need for further research on the issue. However, this result is not unique for Colombia.

At the beginning of the 1980's, Meese and Rogoff (1983a, 1983b) demonstrated the low power of the theoretical models –available at the time– to explain, in a systematic way, the behavior of the main nominal exchange rates. They also showed the low power these models had in predicting the exchange rates. At the beginning, their results looked awkward, but later on several researchers found evidence in the same direction. (Frenkel and Rose (1995) present a review of the main attempts –failures– to overturn the Meese and Rogoff results).

These frustrating conclusions motivated the development of different theoretical models during the 1990's. But, despite these theoretical developments, this new set of models has not been able to provide a great advance in the understanding of the main exchange rates' movements. (See for example Cheung, Chinn and Garcia (2003).

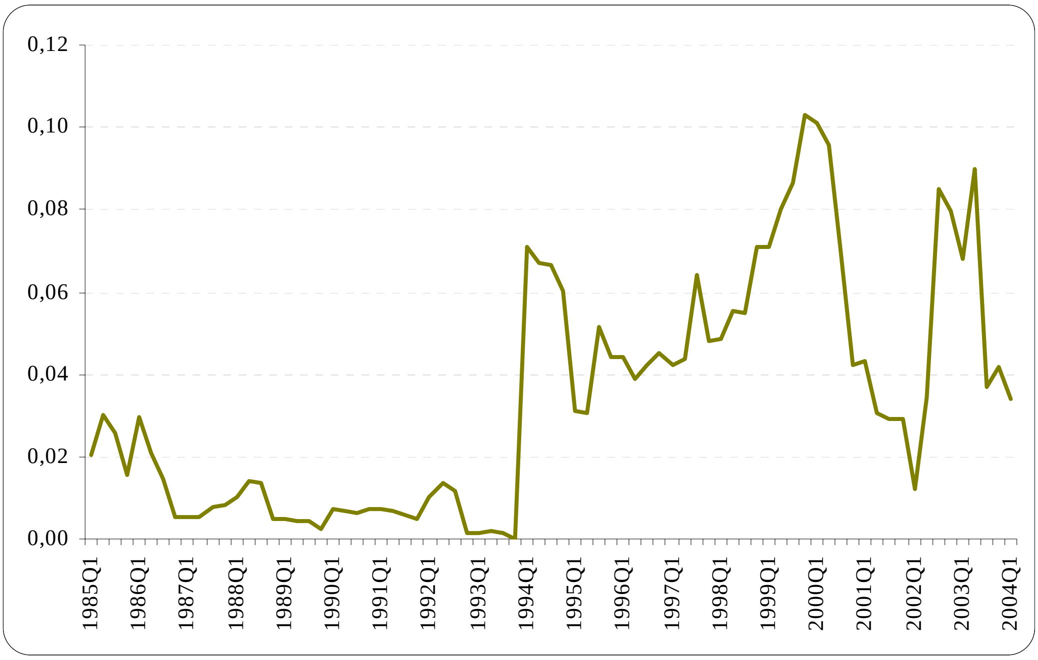
This paper analyzes the out-of-sample forecasting performance of several theoretical models for the USD/COP nominal exchange rate during the period 1984:I-2004:I. During this particular period, Colombia has had different exchange rate regimes (for an extensive discussion on the Colombian exchange rate regimes, see Villar and Rincón (2001), Cárdenas (1997), and Alonso and Cabrera (2004)).

The crawling peg system, which was adopted in 1967, was used until 1991. Under this regime, the Colombian peso was pegged to the US dollar. The pre-specified exchange rate was devalued daily at a continuous devaluation rate. Furthermore, this regime was combined with a system of capital controls. A modified regime was adopted from 1991 until January, it allowed for more flexibility thanks to an implicit exchange rate band. During this period, a foreign exchange market was created, but the Banco de la República (Central Bank) continued to intervene in the market.

In January 1994 an official crawling band system was adopted. Under this regime, the exchange rate was allowed to fluctuate around a pre-determined rate. This band was in use until 1999 when a freely floating system was adopted. Under this free floating regime, the central bank is allowed to reduce the short-run volatility of the exchange rate under pre-established conditions. If the average exchange rate of any day is 5% above or under its 20-day moving average, the central bank may intervene in order to target its volatility. The new regime has induced a change in the volatility of the exchange rate, especially after 1994. Graph 1 shows the volatility of the exchange rate measured as the standard deviation of the first differences of the log nominal exchange rate (USD/COP).

[1] According to the accepted notation, USD/COP stands for the rate of exchange between the United States and Colombia expressed as Colombian pesos per US dollars.

Graph 1: Volatility of the first differences of log USD/COP nominal exchange rate.



Note: Calculated by a rolling standard deviation with a 4 period moving window.
Source: IMF-International Financial Statistics and own calculations.

As noted before, the recent movements of the USD/COP nominal exchange rate do not appear to be convincingly explained by the conventional models adopted in the 1970's. A brief review of the literature about the nominal exchange rate fundamentals may show several approaches. For instance, Cárdenas (1997) evaluates the empirical validity of models that are based on the modern assets approach. According to this approach, the nominal exchange rate fluctuates in order to equilibrate the external demand for domestic assets. For his purpose, Cárdenas (1997) adopts a simple monetary model with flexible prices, a monetary model with sticky prices and a portfolio balance model in order to evaluate the theoretical consistency of such models. In his work, the forecasting performance of such models is not evaluated.

On the other hand, accounting for the main exchange rates, Meese and Rogoff (1983a) analyze the out-of-sample forecasting performance of several structural models for the nominal exchange rate. Their results show how the existing empirical models fail to outperform a random walk. The flexible and sticky price monetary models have a poor performance; despite they base their predictions on the current values of the independent variables.

More recently, Cheung *et al* (2003) argue that the dollar and euro fluctuations can not be explained by the traditional models proposed during the 1970's. They point out that the recent exchange rate movements can only be explained by empirical and theoretical results related with the correlation between variables such as the positions of external assets and the real exchange rate as well as the productivity differentials. They also argue that the forecasting performance of this new set of models (proposed during the 1990's) have not been systematically evaluated. Following their arguments, in order to measure the forecasting performance, a set of metrics is employed by de authors: Mean Squared Error (MSE), direction-of-change and the consistency test developed by Cheung and Chinn (1998). Their results show that according to the MSE metric, none of the models outperforms the random walk. However, according to the direction-of-change metric, some structural models do outperform the random walk.

Taking their work into account, this document analyzes three of the specifications proposed by them for the USD/COP nominal exchange rate. Employing quarterly data for the out-of-sample forecasts, this document relies on four metrics: the Root Mean Squared Error (RMS), the Mean Absolute Error (MAE), the Mean Absolute Percentage Error (MAPE) and the U-Theil.

This document is divided in six sections including this introductory section. The second section presents the theoretical models to be evaluated. The third section analyzes the properties of the series that are employed in the present analysis (Unit Root Tests). The fourth section shows if there is a long-run relation (cointegration analysis) between the variables for each specification. The fifth section presents the forecasting performance metrics for each one of the specifications. Finally, the last section presents some comments on the results.

II. THEORETICAL MODELS TO BE EVALUATED

There are a good number of theoretical models that have been used to model the behavior of the nominal exchange rate (Chinn (1997), Dornbusch (1976), Cárdenas (1997), Owen (2001), Rosenberg (2000) and Meese and Rogoff (1983) among others). In order to analyze the Colombian case, the models proposed by Cheung *et al* (2003) are adopted. First of all, a random walk is defined as a benchmark model. Additionally, as a reference specification, the PPP condition is employed. Following the latter, the identity is defined as:

$$s_t = \beta_0 + \hat{p}_t \qquad (1)$$

where s_t represents the log nominal exchange rate (USD/COP), p_t the log price level (CPI with 2000 = 100) and "^" denotes the intercountry difference. Given that (1) is an identity, it is not estimated.

In second place, the monetary model with sticky prices is considered. In this model the PPP condition is omitted because it is assumed that the short-run prices are sticky. Following Dornbusch (1976), an increase in the domestic monetary supply causes a capital outflow (due to the fall in the interest rate) which causes the exchange rate to depreciate. This model can be expressed as:

$$s_t = \beta_0 + \beta_1\hat{m}_t + \beta_2\hat{y}_t + \beta_3\hat{i}_t + \beta_4\hat{\pi}_t + \varepsilon_t \qquad (2)$$

where m_t, y_t, i_t and π_t denote the log of money supply (M1), real GDP (at 2000 prices), interest rate, and inflation rate, respectively.

Equation (2) can be interpreted as an extension of equation (1) where the macroeconomic variables capture the demand for money and the overshooting effects. Although Cárdenas (1997) imposes restrictions on the coefficients from equation (2), Cheung *et al* (2003) suggest that theory does not present a clear guide to do so. Therefore, equation (2) is estimated without imposing restrictions on the coefficients.

The next two models come from the Balassa-Samuelson (BS) approach. This kind of models accords a central role to productivity differentials in order to explain nominal exchange rate variations. Specifically, Clements and Frenkel (1980) and Chinn (1997) employ productivity differentials to explain the exchange rate. A specification that includes the productivity differential can be expressed as:

$$s_t = \beta_0 + \beta_1 \hat{m}_t + \beta_2 \hat{y}_t + \beta_3 \hat{i}_t + \beta_5 \hat{z}_t + \varepsilon_t \quad (3)$$

where z_t represents a measure of productivity (index of production per employee). Unlike equation (2), this equation does not assume that the PPP holds in the long-run. Therefore, the inflation rate differential is not included. In this kind of models, the exchange rate path is determined by the productivity differentials.

Finally, the fourth model incorporates the BS effect via productivity differentials, the portfolio balance effect via variables such as the ratio between debt and GDP and net foreign assets and, the overshooting effect via real interest rate. This "composite" model can be expressed as:

$$s_t = \beta_0 + \hat{p}_t + \beta_5 \hat{z}_t + \beta_6 \hat{r}_t + \beta_7 \hat{g}debt_t + \beta_8 tot_t + \beta_9 nfa_t + \varepsilon_t \quad (4)$$

where r_t is the real interest rate, $gdebt_t$ represents the government debt to GDP ratio, tot_t the log terms of trade, and nfa_t the net foreign assets. A unity coefficient is imposed on the intercountry log price level $(\hat{p}_t)$. Therefore, (4) can be re-expressed in order to determine the real exchange rate.

III. SERIES AND UNIT ROOTS

Quarterly data ranging from the first quarter of 1984 to the first quarter of 2004 for Colombia and the United States are employed. Most of the series are obtained from the IMF International Financial Statistics Database. The following series are employed: the period average nominal exchange rate (USD/COP); CPI with 2000 as base year; M1 is used as a proxy for the monetary supply (in the case of Colombia, it is obtained from the Banco de la República database); the three-month interests rates are the prime rate and the deposit rate for the Colombian case (the latter from the National Planning Department historical statistics); the terms of trade are calculated from the export and import price indexes; the net foreign assets for Colombia are from the National Planning Department historical statistics; and the real GDP at 2000 prices are from the National Accounts published by DANE and the Bureau of Economic Analysis (NIPA Tables). The series are not seasonally adjusted.

The productivity series is calculated dividing the real GDP by the number of employees and taking it into an index with 2000 as base year. The employment series for Colombia is obtained from Lasso (2002). The employment series for the United States is obtained from the Bureau of Labor Statistics. Both series are not seasonally adjusted. Finally, the government debt to GDP ratio is calculated employing a proxy for debt. Given the data unavailability, it is necessary to employ the Financing series published by the IMF-IFS.

In order to determine the properties of the series (integration order), their stationarity is analyzed through several tests. The Augmented Dickey-Fuller (1979), the Phillips-Perron (1988) and the non-parametric Breitung (2002) tests are initially employed. Additionally, the unit root test proposed by Kwiatkowski, Phillips, Schmidt and Shin (1992) (KPSS) is applied. The results are reported in Table A-1 at the end of the document. The tests are applied to both the levels and the first differences of all series. According to the reported results, they provide some evidence to assert that all the series except $\hat{z}_t$ and tot_t are I(1). For the case of the productivity differential and terms of trade, the evidence shows that they may be I(0).

To obtain robust evidence, further checking is needed. Therefore, additional testing is applied to some of the series. According to the particular characteristics of the series, the unit root test provided by Perron (1989) is carried out to the series that show some kind of structural break. Model A is employed for the series $\hat{y}_t$, $\hat{z}_t$, $\hat{\pi}_i$ and $\hat{i}_t$ given that an exogenous change is observed in their levels. For the series $\hat{p}_t$, given that it shows an exogenous change in its growth rate, the Model B is employed. The results are reported in Table A-2. As observed, the t-statistics are very high and therefore it is not possible to reject the null of a unit root. Summarizing, this test provides some extra evidence about the integration order (I(1)).

One last test is carried out. The non-linear trend stationarity test provided by Bierens (1997) is applied to the series that may show a non-linear trend: $\hat{p}_t$, nfa_t and $\hat{g}debt_t$. The null hypothesis of this test implies the presence of a unit root with drift. The advantage of this test relative to the Perron (1989) test is that it is not necessary to determine the break point (reference date) because when including the polynomial, the data is smoothed. The results are reported in Table A-3. According to the reported results, the statistics for each case point out that $\hat{p}_t$ has a unit root with drift (comparing the four statistics, the null can not be rejected). On the other hand, the results are not conclusive for the case of nfa_t which corroborates the previous results obtained from the conventional tests. In relation to $\hat{g}debt_t$, the results point out that there is no unit root but again, they are inconclusive to determining the non-linearity of the trend.

In summary, according to the conventional tests and the special cases as structural break and non-linear trend tests, it can be asserted that there is enough evidence to assume that the series are integrated with order one.

IV. LONG-RUN RELATION

In order to determine whether there is a long-run relation and discard possible spurious relations, the Johansen Cointegration test (1988) is used. The results are reported in Table 1. According to the reported statistics, for the equation (2) there are two cointegration vectors, three for equation (3) and one vector for equation (4). Taking these results into account it is possible to estimate the proposed specifications using OLS. In the next section, the rolling estimation methodology for obtaining out-of-sample forecasts is discussed.

Table 1: Cointegration test for each structural specification. 1984:I-2004:I.

Johansen Cointegration Test

Statistics for each hypothesis

H_O	H_A	λ – max		
		Equation (2)	Equation (3)	Equation (4)
r = 0	r = 1	52,6 **	105,6 **	52,0 **
r = 1	r = 2	32,3 **	45,7 **	26,3
r = 2	r = 3	15,6	26,4 **	20,2
r = 3	r = 4	6,9	11,1	16,2
r = 4	r = 5	1,3	3,4	8,7
r = 5	r = 6	---	---	0,1
		Trace		
		Equation (2)	Equation (3)	Equation (4)
r <= 4	r = 5	1,3	3,4	---
r <= 3	r = 5	8,2	14,5	---
r <= 2	r = 5	23,8	41 **	---
r <= 1	r = 5	56,1 **	86,6 **	---
r <= 0	r = 5	108,7 **	192,3 **	---
r <= 5	r = 6	---	---	0,14337
r <= 4	r = 6	---	---	8,7976
r <= 3	r = 6	---	---	24,953
r <= 2	r = 6	---	---	45,133
r <= 1	r = 6	---	---	71,481 **
r <= 0	r = 6	---	---	123,47 **

(**): Rejects null hypothesis at a 5% significance level.

V. FORECASTING PERFORMANCE

For estimation purposes, this document adopts the methodology established by Meese and Rogoff. This methodology has been widely implemented in the empirical exchange rate modeling and forecasting literature and it consists in applying rolling regressions. Starting with a given data sample, the estimates are obtained, out-of-sample forecasts produced and then the sample is moved up (rolled) forward one period. The process is repeated until the total sample is worn out.

In this document, 50 percent of the total sample is left for the out-of-sample forecasts, that is, the rolling procedure starts at the middle of the full sample and then begins to move up until the sample is exhausted and the last forecast is produced. It is worth noting that for each one of the specifications, one, two, three and four step-ahead forecasts are computed.

According to Cheung *et al* (2003), rolling regressions do not represent significative efficiency gains but they have the benefit of accounting for parameter instability, a problem that is very common when modeling exchange rates.

To evaluate the forecasting accuracy four metrics are used. The root mean squared error (RMS), the mean absolute error (MAE), the mean absolute percentage error (MAPE) and the Theil's inequality indicator (U-Theil). The first two metrics measure the average difference between the observed values and the forecasted values. The third metric measures the same difference but in relative terms. The four metrics can be defined as:

$$RMS = \sqrt{\frac{1}{T}\sum_{t=1}^{T}\left(Y_t^S - Y_t^a\right)^2}$$

$$MAE = \frac{1}{T}\sum_{t=1}^{T}\left|\left(Y_t^S - Y_t^a\right)\right|$$

$$MAPE = \frac{1}{T}\sum_{t=1}^{T}\left|\frac{Y_t^S - Y_t^a}{Y_t^a}\right|$$

$$U = \frac{\sqrt{\frac{1}{T}\sum\left(Y_t^S - Y_t^a\right)^2}}{\sqrt{\frac{1}{T}\sum_{t=1}^{T}\left(Y_t^S\right)^2} + \sqrt{\frac{1}{T}\sum_{t=1}^{T}\left(Y_t^a\right)^2}}$$

where Y_t^S , Y_t^a and T represents the forecasted value at horizon *h,* the observed value at time *t* and the number of effective out-of-sample forecasts respectively. Table 2 reports the results for each one of the structural models and also for the reference specifications, the random walk and the PPP condition.

The results imply that no model can outperform the random walk at any of the forecasting horizons accounted. Leaving aside the random walk and focusing only on the three structural models and the reference PPP identity leads to different conclusions. According to the RMS metric, at all the four horizons the best specification is the reference PPP condition. When taking into account the MAE and MAPE metrics, the PPP condition outperforms the structural models in forecasting at three and four horizons, but equation (2), the sticky price model, outperforms the rest of the specifications at one and two horizons. That is, the PPP condition can forecast the long-run values better than the rest of the structural models and when forecasting the short-run values, the sticky price model performs better. Finally, according to the Theil´s inequality coefficient, the results are identical to those found with the RMS metric. This result is expected because the U-Theil coefficient measures the error RMS in relative terms.

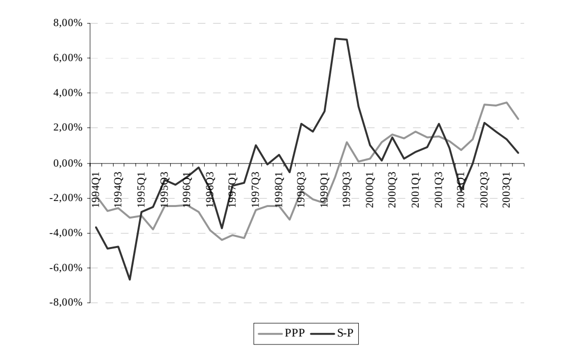
Table 2: Forecasting performance metrics.

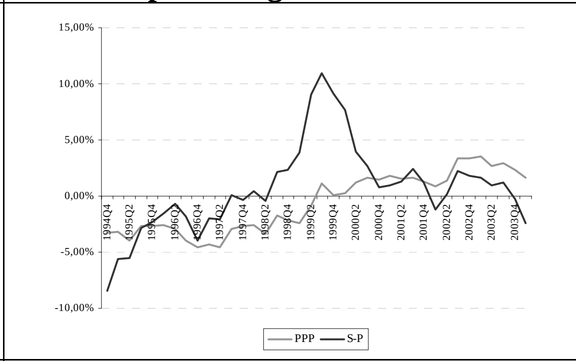
	RMS	MAE	MAPE	U-Theil
PPP (1)				
1 quarter	0,1837	0,1673	0,0231	0,0125
2 quarters	0,1893	0,1729	0,0238	0,0128
3 quarters	0,1922	0,1755	0,0241	0,0130
4 quarters	0,1940	0,1768	0,0241	0,0130
S-P (2)				
1 quarter	0,2002	0,1482	0,0205	0,0137
2 quarters	0,2366	0,1706	0,0235	0,0161
3 quarters	0,2657	0,1895	0,0260	0,0180
4 quarters	0,2923	0,2100	0,0286	0,0198
B-S (3)				
1 quarter	0,2184	0,1632	0,0232	0,0148
2 quarters	0,2382	0,1751	0,0249	0,0161
3 quarters	0,2552	0,1901	0,0268	0,0171
4 quarters	0,2662	0,2027	0,0284	0,0178
Composite (4)				
1 quarter	0,1926	0,1647	0,0217	0,0129
2 quarters	0,2109	0,1817	0,0239	0,0141
3 quarters	0,2378	0,2054	0,0270	0,0159
4 quarters	0,2537	0,2185	0,0287	0,0170
RW (5)				
1 quarter	0,0584	0,0477	0,0065	0,0061
2 quarters	0,0577	0,0466	0,0063	0,0060
3 quarters	0,0591	0,0481	0,0065	0,0061
4 quarters	0,0596	0,0485	0,0065	0,0061

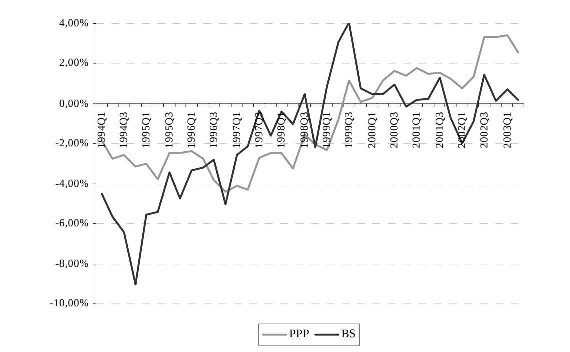
Now, focusing into the graphic analysis, Graph 2 presents the forecasting percentage errors of the equations (2), (3) and (4) for the short and long horizons forecasts compared against the forecasts produced by the PPP condition. As seen on the graph, the forecast percentage error presents a large increase during 1999 coinciding with the adoption of the freely floating exchange system. The graphs also show that the "composite" model is the one with the poorest performance. In relation to this "composite" model, it is worth noting that this specification may present some measurement problems related to the employment of the financing series as a proxy for government debt and also the net foreign assets series. These problems may influence the forecasting errors for this model in a significative way. Therefore, reducing the set of models to the sticky price (S-P) and the Balassa-Samuelson, it can be found that the forecasted percentage errors for the PPP condition present a behavior that appear to be less variable than the behavior of the percentage errors obtained for the S-P and BS models. Also, referring to this last point, the PPP forecasting percentage errors do not present the large increase that can be observed in the rest of the specifications. In conclusion, the graphic analysis may serve as a support for the previous statistical analysis.

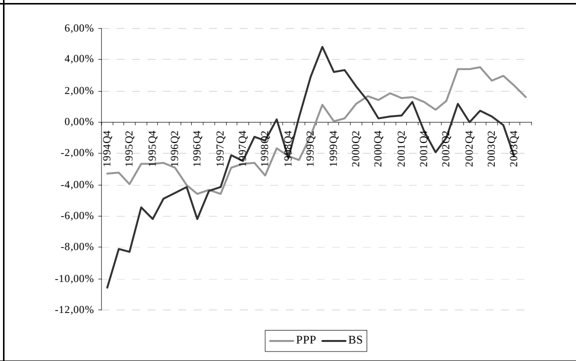
Graph 2: Forecast percentage errors for the structural models.

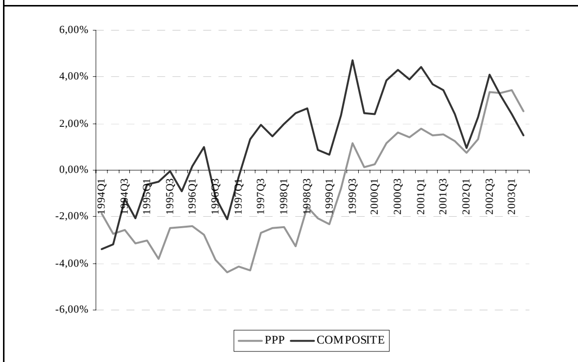
Forecast percentage error at 1 horizon	Forecast percentage error at 4 horizons

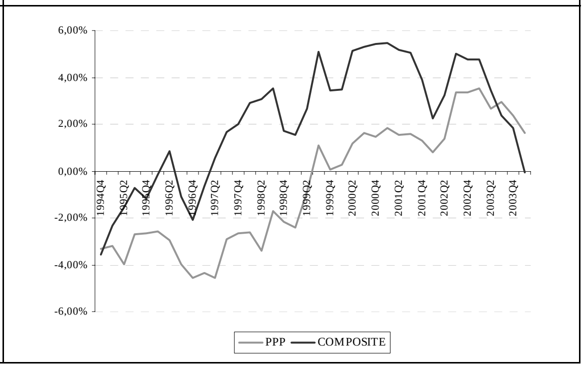












VI. FINAL COMMENTS

This document analyzes the forecasting performance of three models for the USD/COP nominal exchange rate during the period ranging from the first quarter of 1984 to the first quarter of 2004. The models correspond to the monetary and productivity differentials approaches and a “composite” model that includes the BS effect and the portfolio balance effect. The three specifications proposed are compared against a random walk (benchmark model) and the PPP condition which is used as a reference specification.

Initially, the properties of the series are analyzed finding evidence that suggest they are integrated of order one. For this purpose, conventional tests for unit roots and for some special cases, additional tests are employed. The next step is to determine whether they are cointegrated in each one of the specifications proposed. For this purpose, the Johansen (1988) procedure is employed. According to the cointegration test, there is strong evidence about the long-run relation among the variables for each specification and therefore they can be estimated using OLS methodology.

Finally, through several metrics, the forecasting performance of each structural model is evaluated and compared against a naïve random walk. The results suggest that the theoretical models developed during the 1990 are not able to outperform the random walk in producing out-of-sample forecast for one to four quarters ahead. However, comparing the metrics against the reference model (PPP condition), the results suggest that for short-run forecasts (one and two quarters ahead) the sticky price model appears to outperform the PPP condition (MAE and MAPE metrics). As argued by Cheung *et al* (2003), these results are not a clear indicator about the poor forecasting performance of the analyzed structural models. That is, although the results strengthen the previous results obtained by Cheung *et al* (2003), this exercise does not involve all the available metrics that can be used to evaluate the forecasting performance. In addition, the metrics that are employed in this document rely on the assumption of a symmetric loss function. In relation to the loss function, it is worth noting that it would be better to assume an asymmetric loss function given the nature of the exchange markets.

As a final conclusion, it is well known that the exchange markets are a clear example of a market with imperfect information and high transaction costs. This characteristic may lead to the need of implementing non-linear specifications or estimation methods in order to capture possible asymmetries that could be present in the exchange markets. As previously noted in the introductory section, the Meese and Rogoff (1983a, 1983b) results have not been overturned despite the longer data-sets available, and the developing of more advanced econometric techniques. Kilian and Taylor (2001) argue that the explanation professional exchange rate forecasters have given to the issue –that the simple standard economic models are inadequate– is not totally convincing. Instead, Kilian and Taylor (2001) suggest that there is a deep work which documents the various non-linearities in deviations of the spot exchange rate from economic fundamentals[2]. Therefore, they argue that the poor forecast performance can be explained by the fact that despite the sound theory, the empirical implementation as a linear statistical model presents a lot of flaws. Summarizing, with an appropriate non-linear structure, economic models of the exchange rate may prove useful for forecasting at least at longer horizons.

REFERENCES

Alonso, J.C., and A. Cabrera (2004) "La Tasa de Cambio Nominal en Colombia" *Apuntes de Economia – Universidad ICESI*, No2, Junio.

Balke, N.S., and T.B. Fomby (1997): "Threshold Cointegration" *International Economic Review*, 38: 627-45.

Bierens, H. J. (1997): "Testing the Unit Root with Drift Hypothesis against Nonlinear Trend Stationarity, with an Application to the US Price Level and Interest Rate" *Journal of Econometrics* 81: 29-64.

Breitung, J. (2002): "Nonparametric Tests for Unit Roots and Cointegration" *Journal of Econometrics* 108: 343-63.

[2] Balke and Fomby (1997), Taylor and Peel (2000) and Taylor, Peel and Sarno (2001)

Cárdenas, M. (1997): “La tasa de cambio en Colombia” *Cuadernos de Fedesarrollo 1*, Tercer Mundo Ed., Bogotá. Septiembre.

Cheung, Y., M. Chinn and A. Garcia. (2003): “Empirical Exchange Rates of the Nineties: Are Any Fit to Survive?” *Santa Cruz Center for International Economics* Paper 03’14.

Cheung, Y. and M. Chinn. (1998): “Integration, Cointegration, and the Forecast Consistency of Structural Exchange Rate Models” *Journal of International Money and Finance* 17: 813-30.

Chinn, M. (1997): “Paper Pushers or Paper Money? Empirical Assessment of Fiscal and Monetary Models of Exchange Rate Determination” *Journal of Policy Modeling* 19:51-78.

Clements, K. and J. Frenkel. (1980): “Exchange Rates, Money, and Relative Prices: The Dollar-Pound in the 1920’s” *Journal of International Economics* 10: 249:62.

Dickey, D.A. and Fuller, W.A. (1979): “Distribution of the Estimators for Autoregression Time Series with a Unit Root”. *Journal of the American Statistical Association,* 74: 427-31.

Dornbusch, R. (1976): “Expectations and Exchange Rate Dynamics” *Journal of Political Economy* 84: 1161-76.

Engle, R.F. and Granger, C.W. (1987): “Cointegration and Error Correction: Representation, Estimation, and Testing”. *Econometrica* 55: 251-76.

Frankel, J. A. (1979): “On the Mark: A Theory of Floating Exchange Rates Based on Real Interest Differentials” *American Economic Review* 69: 610-22.

Frenkel, J. A. y Rose, A. (1995): “Empirical Research on nominal Exchange Rate” in Grossman and Rogoff: Handbook of International Economics: 1689-729

Hylleberg, S., R. F. Engle, C. W. J. Granger, and B. S. Yoo. (1990): “Seasonal Integration and Cointegration” *Journal of Econometrics* 44: 215-38.

Johansen, S. (1988): “Statistical Analysis of Cointegrating Vectors”. *Journal of Economic Dynamics and Control* 12: 231-54.

Kilian, L. and M.P. Taylor (2001): “Why is it so Difficult to Beat the Random Walk Forecast of Exchange Rates?” *Tinbergen Institute Discussion Paper* No 031-4, January.

Kwiatkowski, D., P. Phillips, P. Schmidt, and Y. Shin. (1992): “Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure Are We that Economic Time Series Have a Unit Root?” *Journal of Econometrics* 54: 159-78.

Lasso, F. J. (2002): “Nueva metodología de la Encuesta de hogares ¿Más o menos desempleados?” DNP, *Archivos de Economía,* Doc 213.

MacKinnon, J.G. (1991): “Critical Values for Cointegration Tests”. In *“Long-run Economic Relationships: Readings in Cointegration”*, (Eds) R.F. Engle and C.W.J. Granger, Oxford University Press, Oxford, pp 267.

Meese, R. A. and K. Rogoff. (1983a): “Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample?” *Journal of International Economics* 14: 3-24.

Meese, R. A. and K. Rogoff. (1983b): “The out-of-sample failure of empirical exchange rate models: sampling error or misspecification?” en Exchange Rates ad International Macroeconomics: NBER y University of Chicago Press: 67-105.

Owen, D. (2001): “Importance of Productivity trends for the euro” *European Economics for Investors*, March.

Perron, P. (1989): “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis” *Econometrica* 57: 1361-401.

Pindyck, R.S y D. L. Rubinfeld. (1998): “Econometría: Modelos y Pronósticos” McGraw – Hill Ed., México. Cuarta Edición.

Phillips, P. and Perron, P. (1988): “Testing for a Unit Root in Time Series Regression” *Biometrika* 75: 335-46.

Rosemberg, M. (2000): “The Euro’s Long-Term Struggle” *FX Research Special Report Series*, No 2.

Taylor, M.P. and D.A. Peel (2000): “Nonlinear Adjustment, Long Run Equilibrium and Exchange Rate Fundamentals” *Journal of International Money and Finance,* 19: 33-53.

Taylor, M.P., D.A. Peel, and L. Sarno (2001): “Nonlinear Adjustment in Real Exchange Rates: Towards a Solution to the Purchasing Power Parity Puzzles” forthcoming: *International Economic Review*.

Villar, L. and H. Rincón (2001): “Flujos de capital y regímenes cambiarios en la década de los 90” *Revista ESPE, Banco de la República.* No 39, Junio.

TABLE A-1

Conventional Unit Roots Tests

Statistics for each test.

		Levels /1				*First Differences /2*			
		ADF /3	**PP**	**Breitung (2002) /4**	**KPSS**	**ADF**	**PP**	**Breitung (2002) /4**	**KPSS**
s_t		-2,000	-4,150	0,017	0,147 ++	-2,559 /5	-91,310 ººº	0,013	0,443
$\hat{p}_t$		-1,168	2,480	0,022	0,163 ++	0,019	-88,460 ººº	0,031	0,478 ++
$\hat{m}_t$		-0,916	-2,860	0,021	0,156 ++	-1,384 /6	-75,510 ººº	0,003 ººº	0,254
$\hat{y}_t$		-0,698	-25,760 ºº	0,018	0,146 +	-2,494 /6	-116,220 ººº	0,001 ººº	0,145
$\hat{\pi}_i$	*NT*	-0,516	-1,630	0,060	0,407 +	-2,067 /5	-54,220 ººº	0,005 ººº	0,382 +
$\hat{i}_t$		-3,583 ºº	-9,150	0,010	0,152 ++	-2,387 /5	-33,550 ººº	0,001 ººº	0,200
$\hat{z}_t$		-1,894	-75,090 ººº	0,003	0,074	-2,837 º	-102,870 ººº	0,000 ººº	0,087
$\hat{r}_t$	*NT*	-1,599	-11,690	0,040	0,510 ++	-2,664 º	-41,020 ººº	0,000 ººº	0,106
$\hat{g}debt_t$	*NT*	-1,215	-122,370 ººº	0,049	0,463 +	-2,927 ºº	-129,010 ººº	0,000 ººº	0,128
tot_t	*NT*	-2,685 °	-11,710	0,013 º	0,140	-1,892 /5	-42,770 ººº	0,001 ººº	0,130
nfa_i		-0,886	-1,110	0,021	0,170 ++	-3,103 ºº	-92,320 ººº	0,014	0,461 +
$s_t - \hat{p}_t$		-2,097	-8,870	0,011	0,092	-2,939 °°	-81,710 ººº	0,003 °°°	0,148

ADF, PP and Breitung (2002): Correspond to the statistics for the stationarity tests: Dickey-Fuller, Phillips-Perron and Breitung (2002).
KPSS: Corresponds to the statistic for the Kwiatkowski, Phillips, Scmidt y Shin (1992) unit root test.
(º): Rejects the null of a process with a unit root at a 10% significance level.
(ºº): Rejects the null of a process with a unit root at a 5% significance level.
(ººº): Rejects the null of a process with a unit root at a 1% significance level.
(+): Rejects the null of a stationary process around a trend at a 10% significance level.
(++): Rejects the null of a stationary process around a trend at a 5% significance level.
/1: Unless otherwise stated, all cases include a trend given that most of the series present a growth rate.
/2: A trend is not included in the cases for the firts differences.
/3: The optimal number of lags is determined according to the Akaike Information Criteria (AIC).
/4: The decision is based on the critical values simulated from 1000 iterations for a Gaussian process.
/5: When carrying the test according to the number of lags suggested by the SBC criteria, the null of a process with a unit root is rejected at a 1% significance level.
/6: When carrying the test according to the number of lags suggested by the SBC criteria, the null of a process with a unit root is rejected at a 5% significance level.
NT: A trend is not included for the test on the levels.

TABLE A-2

Unit Root Test with Structural Break - Perron (1989)

t statistics.

	Reference date	Case /1	t statistic	Lag number /2
$\hat{y}_t$	1998:IV	A	8,93E+09	6
$\hat{z}_t$	1990:IV	A	1,22E+14	8
$\hat{\pi}_i$	1999:I	A	3,98E+13	8
$\hat{i}_t$	1999:I	A	2,14E+13	9
$\hat{p}_t$	1998:IV	B	2,37E+11	5

/1: Model A represents the Crash Model (structural break). Model B represents the Changing Growth Model (Exogenous change on the growth rate).
/2: Determined employing the Akaike Information Criteria (AIC).
Note: In none of the cases it is possible to reject the null of a unit root.

TABLE A-3

Unit Root Test (Non-Linear Trend) - Bierens (1997)

Statistics for each case (p values in parenthesis) /1

$\hat{p}_t$		*t(m)*	*A(m)*	*F(m)*	*T~(m)*
p = 5	m = 8	-2,791	-59,466	4,221	5289,837
		(0,4210)	(0,2560)	(0,4310)	(0,7080)
	m = 9	-3,565	-138,021	5,002	5369,011
		(0,4080)	(0,1740)	(0,4940)	(0,5700)
	m = 10	-3,661	-153,623	4,597	5548,085
		(0,3480)	(0,1370)	(0,4160)	(0,5770)
nfa_i					
p = 0	m = 8	-6,142 º	-53,649	7,189 **	2006,201 ***
		(0,0990)	(0,1190)	(0,9710)	(0,9910)
	m = 9	-5,575	-51,271	6,521 *	2298,907 **
		(0,3130)	(0,2850)	(0,9440)	(0,9770)
	m = 10	-5,536	-51,285	5,844	2373,139 **
		(0,4520)	(0,3820)	(0,8840)	(0,9600)
$\hat{g}debt_t$					
p = 1	m = 7	-8,140 ººº	-151,986 ººº	8,822 ***	65,809
		(0,0000)	(0,0090)	(0,9990)	(0,1900)
	m = 8	-8,376 ººº	-163,933 ººº	8,286 ***	66,378 º
		(0,0000)	(0,0000)	(1,0000)	(0,0890)
	m = 9	-8,327 ººº	-165,632 ººº	7,384 ***	79,990 º
		(0,0000)	(0,0000)	(0,9980)	(0,0710)

p: Number of lags included in the auxiliary equation (AIC).

m: Polynomial order.

/1: The decision is based on critical values simulated from 1000 iterations.

(º): Rejects to the left at a 10% significance level

(ºº): Rejects to the left at a 5% significance level

(ººº): Rejects to the left at a 1% significance level

(*): Rejects to the right at a 10% significance level

(**): Rejects to the right at a 5% significance level

(***): Rejects to the right at a 1% significance level